The Effects of Liberalization on Market and Currency Risk in the European Union

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Abstract

This paper investigates the effects of liberalization on the pricing of market and currency risk for a number of financial markets in the European Union (EU). An International Asset Pricing Model with a multivariate GARCH-in-Mean specification and time-varying prices of risk is used for the four markets with the largest market capitalization in the EU. Only one price of market risk exists and international investors are rewarded for their exposure to currency risk. The evidence shows that all prices of risk are time-varying and have been decreasing during the process of liberalization. There is also evidence that financial markets react to periods of uncertainty in the process toward the completion of liberalization. In addition, the operation of the European Monetary System has generated lower covariances. As a consequence, total risk premia have declined in the last decade.

JEL Classification: G12, G15

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1. Introduction

This paper studies the impact of liberalization in European financial markets. In particular, it addresses the important issue of whether relaxing restrictions on international investments affects the risk assessment of financial assets.

The institutional and political strivings towards the European Union (EU) have been the driving force behind a number of financial reforms directed at liberalizing European markets. These efforts toward legal integration have raised a number of issues among researchers and policy makers regarding the riskiness of liberalized financial markets, the relevance of currency risk and the need for a single currency.

European markets provide an interesting opportunity to empirically investigate the effects of liberalization and, to my knowledge, this study presents the first attempt. Previous studies on European assets have limited their analysis to the issue of integration, considering either the stock or the money market.[1] Only recently Hardouvelis, Malliaropulos and Priestley [1999] have shown in a conditional framework that stock markets in the member states of the European Monetary Union (EMU) seem to be almost fully integrated. Their paper, however, does not tackle the impact of liberalization. On the other hand, a few papers have recently looked at the issue of liberalization focusing on emerging markets. Henry [2000], Bekaert and Harvey [2000] and Errunza and Miller [2000] start from the assumption that asset markets for emerging economies are segmented and analyze whether changes in asset prices are consistent with a movement toward more integrated markets.

I examine the capital markets of France, Germany, Italy and the United Kingdom in the context of an international asset pricing model derived under integration. Within this framework, I estimate the impact of liberalization on the expected compensation for risk by including conditioning information about the intensity of capital controls and by inferring whether the price of risk was affected in a predictable way.

The asset pricing model originally developed by Adler and Dumas [1983] is used for the analysis. This model offers several appealing features. First, it

[1] This stream of research has analyzed equity and money markets separately, with mostly inconclusive evidence. For the equity market, see for example Beckers, Grinold, Rudd and Stefek [1992], Heston Rouwenhorst and Wessels [1995]. For the money market see Rogoff [1985], Karfakis and Moschos [1990], Katsimbris and Miller [1993], Knot and de Hann [1995], Ayuso and Restoy [1996].

can uncover the relevance of different sources of risk, including currency risk, for different types of assets in an economy. Second, it provides a framework for a conditional test that allows simultaneous investigation of the European stock market and Eurocurrency market.[2] This contrasts with previous unconditional versions of international asset pricing models that could not be supported by the data and in general could not detect the relevance of currency risk.[3] Third, it allows the specification of time-varying prices of risk. This contributes an important feature since it can capture the variation in the compensation for risk resulting from the institutional changes and helps assessing their impact within the asset pricing model.

For simultaneous estimation of the model, I use the parsimonious multivariate GARCH-in-Mean specification proposed by De Santis and Gerard [1997, 1998] because it can be applied to a large number of assets while fully parameterizing risk premia.[4] The presence of GARCH effects has been widely documented in the conditional second moments of all kinds of financial data.[5] Moreover, a fully parametric approach is crucial for the scope of this paper. By explicitly parameterizing prices of risk, I can provide a measure of the impact of liberalization and observe the dynamics of the expected compensation for risk during the progress toward globalization. By also parameterizing second moments, I directly assess the size and economic significance of risk premia and examine how they have changed over time.

My results can be summarized as follows. First, the evidence shows that international investors are uniformly rewarded for their exposure to market and currency risk across European stock and money markets. The size of currency risk is also economically significant over subsamples. Second, prices of risk are time-varying and have been generally falling during the process of markets liber-

[2]See Dumas and Solnik [1995] and De Santis and Gerard [1998] for the only existing tests, with different methodologies, of the conditional version of the model among the four largest world financial markets such as the U.S., Japan, U.K., and Germany.

[3]See Solnik [1974b], Stehle [1977], Korajczyk and Viallet [1989, 1992], Jorion [1991] for examples of unconditional tests of international asset pricing models.

[4]See Bollerslev, Chou and Kroner [1992] for an extensive review of GARCH theory and its applications on stock return, interest rate, exchange rate data.

[5]On European data, GARCH has been successfully applied to currencies (see Bollerslev [1990], Vlaar and Palm [1993], both in a multivariate framework) and equities (see de Jong, Kemna and Kloeh [1992] on Dutch data, Poon and Taylor [1992] on UK stock prices, De Santis [1991] on the Italian market).

alization. In particular, I find that the expected compensation for risk is sensitive to periods of uncertainty related to the success of the European Union, such as the tightening of capital controls or the probability of a currency devaluation. The operation of the European Monetary System (EMS) has also contributed to a decrease in the covariances for currency risk. Overall, total risk premia have declined over time. European financial markets are thus now less risky for investors than they were two decades ago.

Such changes in the reward to risk have also important implications from an international corporate finance perspective. They imply that European firms are facing a lower cost of equity capital which will increase the valuation of cross-border investment projects.

The rest of the paper is organized as follows. Section 2 of the paper illustrates the empirical model. The methodology for estimation is described in section 3. Section 4 presents the data. The empirical results follow in section 5 and section 6 concludes the paper.

2. A Model of International Asset Pricing

Models of international asset pricing stem from the application of the traditional CAPM framework of Sharpe [1964] and Lintner [1965] to an international setting. As it is emphasized in Adler and Dumas [1983], the theory of international finance follows domestic finance theory in deriving equilibrium pricing relationships and risk-return trade-offs from individual portfolio maximization. Their work looks at the optimization problem for investors of different countries in an intertemporal framework.

The theoretical model developed in Adler and Dumas [1983] derives the following equilibrium pricing relation:

$$E[r_{jt} \mid \Im_{t-1}] = \sum_{l=1}^{L} \delta_{l,t-1} cov[r_{jt}, r_{n+l,t} \mid \Im_{t-1}] + \delta_{m,t-1} cov[r_{jt}, r_{mt} \mid \Im_{t-1}] \qquad (2.1)$$

where r_{jt} is the nominal return on an asset or portfolio $j, j = 1, \ldots, N+1$, from time $t-1$ to t, in excess of the risk free rate of the measurement currency, r_{mt}, is the excess return on the world market portfolio (the $(N+1)^{th}$ asset which I denote with the subscript m), and $\Im_{t-1}$ is the information set available to all investors at time $t-1$. The time-varying coefficients $\delta_{l,t-1}, l = 1 \ldots L$, are the world price of exchange rate risk for the L relevant currencies while $\delta_{m,t-1}$ is the world price of market risk. In the theoretical derivation, $\delta_{m,t-1}$ is shown to depend on the relative risk aversion of each group of investors in the economy, weighted by the corresponding relative wealth.

Equation (2.1) represents an innovation with respect to the classic CAPM because of the inclusion of different risk premia. In the traditional CAPM framework, investors are only compensated for their exposure to market risk, which is measured by the covariance between each asset returns and the return on the market portfolio. This implies that $\delta_{m,t-1}$ is a measure of the trade-off between expected return and risk on the market. For this reason, $\delta_{m,t-1}$ is usually referred to as the price of market risk.[6]

In an international framework, it is possible that investors face other forms of risk: that is why equation (2.1) contains a series of additional risk premia. As in the classical case, investors of different nationalities still care about real returns. However, they also face different purchasing power in evaluating returns from the same security and therefore their portfolio composition will differ. If purchasing power parity were to hold, the nominal excess return on a security would depend only on the covariance of the asset with the return on the benchmark portfolio, since there would be no differences in asset appreciation among investors of different countries. However, there is strong empirical evidence that purchasing power deviations are large and persistent over long periods. Therefore international investors seek compensation not only for fluctuations in asset market but also for fluctuations in exchange rates. Even when inflation is considered non-stochastic, exchange rates still show high variability. This additional risk generates a risk premium which depends on the covariances of each asset with the exchange rates. In this sense, the δ_l can be interpreted as the expected compensation for each additional unit of currency risk.

In conclusion, each group will demand a premium for those assets that protect his real purchasing power. The equilibrium relationship in (2.1) is derived aggregating across national groups. As a result, equation (2.1) defines the expected returns in equilibrium as a sum of exchange premia associated with the different

[6]The traditional version of the CAPM from Sharpe [1964] and Lintner [1965] has been extensively estimated, with US data as well as with international data. For this one-factor, international CAPM see, for example, Solnik [1974], Giovannini and Jorion [1989], Engel and Rodrigues [1989], Korajczyk and Viallet [1989], Harvey [1991], Chan, Karolyi and Stulz [1992], De Santis and Gerard [1997].

national groups in the economy. Since exchange rates are highly correlated with purchasing parity deviations, these exchange premia are thus the result of the hedging behavior of investors against shifts in purchasing power.

3. Methodology

In this section I illustrate the methodology developed by De Santis and Gerard [1998] to estimate the version of the conditional asset pricing model described in equation (2.1).

The implications of the conditional version of the model is that optimizing investors update their strategy using the new available information in every period. Conditional versions of the CAPM are usually harder to estimate than unconditional versions. However, the task has been easier since the seminal paper of Engle [1982] on conditionally heteroskedastic processes.

Equation (2.1) has to hold for every asset. In a world with L countries, N risky assets and a world market portfolio, the following system of equations has to hold at every point in time:

$$\begin{array}{rcl}
E[r_{1t} \mid \Im_{t-1}] & = & \sum_{l=1}^{L} \delta_{l,t-1} cov[r_{1t}, r_{n+l,t} \mid \Im_{t-1}] + \delta_{m,t-1} cov[r_{1t}, r_{mt} \mid \Im_{t-1}] \\
 & \vdots & \\
E[r_{n-1,t} \mid \Im_{t-1}] & = & \sum_{l=1}^{L} \delta_{l,t-1} cov[r_{n-1,t}, r_{n+l,t} \mid \Im_{t-1}] + \delta_{m,t-1} cov[r_{n-1,t}, r_{mt} \mid \Im_{t-1}] \\
E[r_{n+1,t} \mid \Im_{t-1}] & = & \sum_{l=1}^{L} \delta_{l,t-1} cov[r_{n+1,t}, r_{n+l,t} \mid \Im_{t-1}] + \delta_{m,t-1} cov[r_{n+1,t}, r_{mt} \mid \Im_{t-1}] \\
 & \vdots & \\
E[r_{n+L,t} \mid \Im_{t-1}] & = & \sum_{l=1}^{L} \delta_{l,t-1} cov[r_{n+L,t}, r_{n+l,t} \mid \Im_{t-1}] + \delta_{m,t-1} cov[r_{n+L,t}, r_{mt} \mid \Im_{t-1}] \\
E[r_{mt} \mid \Im_{t-1}] & = & \sum_{l=1}^{L} \delta_{l,t-1} cov[r_{mt}, r_{n+l,t} \mid \Im_{t-1}] + \delta_{m,t-1} var[r_{mt} \mid \Im_{t-1}]
\end{array} \tag{3.1}$$

The first $n-1$ equations are pricing equity portfolios, the next $L = N-n$ equations are for the pricing of the currency deposits and the last one is pricing the world equity portfolio.

De Santis and Gerard [1998] write equation (2.1) for estimation as:

$$r_t = \sum_{l=1}^{L} \delta_{l,t-1} h_{n+l,t} + \delta_{m,t-1} h_{m,t} + \epsilon_t \qquad \epsilon_t \mid \Im_{t-1} \sim N(0, H_t) \tag{3.2}$$

where r_t is the $s \times 1$ vector of excess returns and H_t is the $s \times s$ conditional covariance matrix whose $(n+l)^{th}$ column is $h_{n+l,t}$ and whose last column is $h_{m,t}$.

This implies that the $(n+l)^{th}$ column of H_t contains the conditional covariances between the returns of each asset and the l^{th}currency deposits. The system of equations in estimation therefore consists of the $1 \leq x < n$ national equity portfolios, L currency deposits and the world market portfolio.

The dynamics of the conditional second moments are left unspecified by the asset pricing model. However, it has been documented in a vast literature that asset prices of both securities and exchange rates exhibit volatility clustering and leptokurtosis. Such characteristics are taken into consideration by GARCH processes and therefore a simple parameterization can offer a testable version of the model.

De Santis and Gerard [1998] specify the dynamics of H_t as:

$$H_t = H_0 * (\iota\iota' - \mathsf{aa}' - \mathsf{bb}') + \mathsf{aa}' * \epsilon_{t-1}\epsilon_{t-1}^{\prime} + \mathsf{bb}' * H_{t-1} \tag{3.3}$$

where a and b are $s \times 1$ vectors of constant coefficients, H_0 is the unconditional variance-covariance matrix of the residuals and * is the Hadamard matrix product (element by element).

The choice of this parameterization is mostly determined by parsimony concerns: in a multivariate system like the one in this study, a general specification for H_t without restrictions would make estimation essentially impossible. Instead, the specification in (3.3) is appealing because it reduces the number of unknown parameters for estimation but still maintains the dynamics of the conditional second moments. I explain in turn each restriction that has been imposed on the proposed parameterization from the general specification.

First, the process is specified as a GARCH(1,1): this is a common simplification in the GARCH literature, since it is often the case that no autocorrelations in the residuals is left unexplained with this lag structure.

Second, the matrices of coefficients characterizing the process are restricted to be diagonal: the diagonal elements of these matrices are the elements of the vectors a and b in (3.3). This implies that each conditional covariance depends only on its own past values and surprises, excluding cross-effects from the other covariances and cross-products of the remaining forecast errors.

In a $s \times s$ system, taking all previous restrictions in consideration, $s(s+1)/2+2s$ parameters have to be estimated only for the covariance equation, which is still a large number. In fact, an estimate for the first term on the right hand side on the right hand side of equation (3.3) is still required, given that H_0 is not directly observable.[7]

Third, Ding and Engle [1994] propose a more parsimonious representation by providing an estimator for the first term. Through the additional assumption of covariance stationarity[8], the sample covariance matrix of the residuals is a consistent estimator of the unconditional covariance matrix. As De Santis and Gerard [1997] argue, the sample covariance matrix of the returns could be used as an approximation for H_0 in estimation, if the conditional mean of the returns were constant. However, in the proposed model under the null hypothesis, the conditional means are time-varying and are a function of second moments, which would make a two-stage estimator of the residuals inconsistent. Therefore they implement an iterative procedure: in the first iteration H_0 is set equal to the sample covariance matrix of the returns, and it is subsequently updated using the covariance matrix of the estimated residuals at the end of each iteration.

In a diagonal system with s assets, the number of unknown parameters in the covariance process is now reduced to $2s$.

In this case, if the assumption of covariance stationarity does not hold over the sample period, inference will be wrong. However, it will be shown that the estimated values in the system imply a covariance stationary process.

Equations (3.2) and (3.3) give the model for estimation.

Assuming a normal conditional density, the log likelihood function is written as

$$\ln L(\theta) = -\frac{Ts}{2}\ln 2\pi - \frac{1}{2}\sum_{t=1}\left[\ln\left|H_{t(\theta)}\right| + \epsilon_t(\theta)' H_t(\theta)^{-1}\epsilon_t(\theta)\right] \tag{3.4}$$

where θ is the vector of unknown parameters in the model. The estimation is performed using the BHHH (Bernt, Hall, Hall and Hausman, [1974]) algorithm. Under regularity conditions, the quasi-maximum likelihood estimator of θ is generally consistent and asymptotically normal, as shown in Bollerslev and Wooldridge [1992]. To avoid incorrect inference due to the mispecification of the conditional density of asset returns, quasi-maximum likelihood estimates for the standard errors are used to guarantee robustness of the results(see White [1982], Bollerslev

[7] In the unconstrained version, the first term on the right hand side is parameterized as a matrix of free elements, $C'C$.

[8] With the assumption of covariance stationarity, the unconditional variance-covariance matrix of the residuals is equal to $H_0 = C'C * (\iota\iota' - \mathrm{aa}' - \mathrm{bb}')^{-1}$ which implies that $C'C = H_0 * (\iota\iota' - \mathrm{aa}' - \mathrm{bb}')$.

and Wooldridge [1992]).[9]

The advantages of this fully parameterized approach are in obtaining estimates of the conditional second moments and of the time-varying risk premia. In fact, it has been claimed that correlations among markets could change as a result of liberalization. Moreover, correlations involving assets in the currency market are likely to be affected by the operation of the EMS. It is thus of primary interest for this study to obtain estimates of risk premia to infer the effects of liberalization. The proposed methodology provides the best approach to analyze these issues.

4. Data and Preliminary Statistics

I analyze the impact of liberalization among four countries: Germany, United Kingdom, France and Italy, which are the countries with the largest market capitalization in the European Union.

This study will take the point of view of a German investor. I will use monthly returns of the countries stock indices and currency deposits over the period March 1974 - August 1995.[10]

The stock market indices are from Morgan Stanley Capital International (MSCI). [11] As a benchmark index, I use a value weighted world index which represents approximately 60% of the aggregate market value of the stock exchanges of twenty countries worldwide. Past studies (see Dumas and Solnik [1995], De Santis and Gerard [1998]) proved capital markets integration for Germany and U.K. with US and Japan by using the world market index. In this paper, a test for integration which instead includes France and Italy will then imply integration with the rest of the world and justifies use of the world index to measure systematic risk.

Currency deposits represent the mark rate of return of a currency holding in country l. The monthly interest rates for these series are from Data Resources Incorporated (DRI). These are the Eurocurrency rates offered in the interbank market in London for 30 day deposits in French franc, Deutsche mark, Italian

[9] Since the likelihood function may not be correct, the robust covariance matrix is equal to $A^{-1}BA^{-1}$, where A^{-1} is the minus the expected value of the Hessian and B is the expected value of the cross-product of the scores.

[10] The choice of the sample period is dictated by data availability.

[11] Use of the MSCI indices is advantageous because they are fully comparable with one another, since they are constructed on the same basis and principles and share an identical base.

lira and British pound. Due to limited availability of data for the Eurolira, I extended this series with short-term interest rates reported by the International Financial Statistics (IFS) and with Eurorates from Bank of International Settlements (BIS).[12]

Monthly returns of both types of assets have been translated in Deutsche mark using the closing European interbank currency rates from MSCI.

The model assumes as a conditionally risk free asset the German short term rate, from DRI.

In total, the system consists of eight equations, one for each of the assets included: four market indices (France, Germany, Italy and United Kingdom), the world index of equity and three currency deposits (French franc, Italian lira and British pound). The pricing equation thus includes three exchange risk premia, one for each of the currencies under consideration.

Panel A of table 1 contains the descriptive statistics for the asset excess returns. As expected, returns from the stock indices show higher mean, as well as higher volatility, than the currency deposits. The history of appreciation characterizing the evolution of the DM vis-a-vis the Italian Lira in the last decade can explain the sizable difference for mean returns of Italian equity from the other equities mean returns. For all the series, the hypothesis of normality is strongly rejected by the Bera-Jarque statistics.

Since the specification of conditional mean and variances is a key issue, I look at the autocorrelations for returns. In general, there is no evidence of significant autocorrelation in the series. The Ljung-Box $Q(z)_{12}$ test statistics in panel A confirm the finding. Conversely, more predominant is the presence of autocorrelations for squared returns in both sets of assets, as revealed from the Ljung-Box $Q(z^2)_{12}$. This is taken as support for the existence of GARCH effects.

The proposed diagonal GARCH parameterization implies that each covariance depends only on its own past values and surprises, excluding cross-effects from the remaining covariances and cross-product of the remaining forecast errors. A potential problem could then be due to cross-market dependencies in volatility. As a check, I calculated the cross-correlations at different leads and lags of the

[12]Line 60B from June 1973 to February 1977, line 60C from February 1977 to August 1977 from IFS, Eurolira rate from September 1977 to November 1980 from BIS. The correlation between the IFS and the DRI series, when overlapping data are available, is .986. The choice of these data for short-term interest rates can be justified after the work of Ferson and Harvey [1993].

squared returns, some of which are reported in panel B. Only 5% of the 112 non simultaneous correlations of order (-2,-1,1,2) between all pairs of assets is statistically significant and not presents a concern.

The main motivation for international diversification relies on correlations among international assets that are lower than correlations among domestic assets. Panel C shows low unconditional correlations among the assets under investigation. Most of the correlations among stock indices are below .5 while the ones among currency deposits are even lower.

The model being tested does not specify state variables that can explain the observed dynamics of the prices of risk. Therefore previous literature[13] utilized as instruments variables that are connected with the evolution of financial markets: these variables are intended to convey the information available to investors. Some of the commonly used instruments are also adopted in this paper. These are the dividend price ratio on the world equity index, the US term premium, the US default premium as the difference between Baa and Aaa rated bonds, the change in the US short-term interest rate. To this list, I add the German term premium and the change in the German short-term interest rates. In fact, these variables could help predict time-variation in the context of financial markets linked to the European Monetary System.

Since my main goal is to evaluate the impact of liberalization, I also need to include conditioning information regarding such process. I therefore look at two alternatives. The first one expands the previous set with a dummy variable with a value of 1 after the deadline set at July 1, 1990 for capital movements liberalization. The second alternative includes in the information set the differentials between the offshore/onshore one-month interest rate of each country.[14] This interest rate differential has been used in the past[15] as an indication of the degree of openness in the capital markets of one country.[16] With tight capital controls,

[13]See for examples Harvey [1991], Ferson and Harvey [1993], Dumas and Solnik [1995], De Santis and Gerard [1997,1998].

[14]I have chosen the short-term differentials with one-month maturity since the most relevant innovation in the liberalization of financial markets has been the abolition of restrictions for short-term capital movements.

[15]Some of these studies used Covered Interest differentials across national boundaries as a measure of capital controls. However, since Covered Interest Parity holds within the Euromarkets, this differential is equivalent to the offshore/onshore differential.

[16]See Dooley and Isard [1980], Claassen and Wyplosz [1982], Rogoff [1985], Giavazzi and Pagano [1988].

onshore and offshore rates can diverge over some period, also allowing countries to pursue independent monetary policies. Conversely, in the absence of legal barriers to international capital movements, onshore and offshore rates move very closely. This variable thus appears best suited to convey information through time about the progress of liberalization.

Summary statistics for the information variables are in panel A of table 2, with the autocorrelations in panel B.[17] As expected, the French and Italian differentials are the ones with the highest mean among the differentials, since these countries maintained capital controls for more than half of the sample. As a check for multicollinearity among all instrumental variables, the unconditional correlations are reported in panel C.

Inspection of the plots of the differentials in figures 1 a through c offers insights on the progress of liberalization. French and Italian differentials follow mostly the same dynamics. After a period of tightened control measures in 1981, both countries started their liberalization process around 1984 and completed it in the first half of 1990, before the deadline of July 1 imposed by the European Union. Equally, a marked change in the dynamics of the U.K. differential corresponds to the abolition of exchange controls for this country in July 1979. Evidence of the change in capital movements regulations can also be found in the respective means. From the beginning of the sample until July 1990, the monthly mean is equal to 14 and 16 basis points for the French and the Italian differential, respectively. It then decreases to 1 and -5 basis points, respectively. For the UK differential, the monthly mean decreased from 18 basis points before liberalization to 7 basis points after the reforms. A zero difference is evidence of no legal barriers within financial markets. This preliminary analysis confirms the validity of the choice of the type of data as instruments to generate time variation in the prices of risk and to infer the impact of liberalization.

5. Empirical Results

This section examines the empirical findings on the expected compensation for risk among the European countries. For my analysis, I proceed in stages and consider a few versions of the benchmark model in (3.2) and (3.3) to establish

[17] The Eurolira one-month rate was not available before November 1977. The series has been filled with rates constructed using covered interest parity with the Eurodollar. Rogoff [1985], for example, uses the same procedure to construct non-dollar Eurorates.

whether the pricing restrictions of each version are satisfied. In doing so, I use the QML Wald statistics suggested by Bollerslev and Wooldridge [1992] which are robust to departure from normality.

5.1. Pricing Restrictions of the model

In this section, I test for the correct pricing restrictions of the model over the whole sample period, without imposing any structural change in the prices of risk. This is a necessary step that will allow me in the following stage to examine the impact of liberalization within the selected asset pricing model specification.

I start by estimating the following version of equation (3.2):

$$r_{jt} = \alpha_j + \sum_{l=1}^{L} \delta_{lt-1} cov_{t-1}[r_{jt}, r_{n+l,t}] + \delta_{mt-1} cov_{t-1}[r_{jt}, r_{mt}] + \epsilon_{jt} \tag{5.1}$$

where α_j are asset specific intercepts, $\delta_{l,t-1}$ are the time-varying prices of exchange rate risk and $\delta_{m,t-1}$ is the time-varying price of market risk. For my dataset, exchange rate premia are the sum of three premia related to the Eurofranc, the Eurolira and the Europound.

The significance of the pricing restrictions of equation (5.1) provides a test of the international version of the CAPM.

First, the theoretical model in equation (3.2) implies that capital markets are integrated. In an asset pricing framework, this means that one model can consistently price all assets and therefore assets with the same risk generate the same expected excess returns, irrespective of where they are traded. Therefore a test of $\alpha_j = 0$ for all j can shed some light on whether European markets are financially integrated.[18] The significance of asset-specific intercepts in this context could be due to difference in taxation among countries, to mild market segmentation, or even to other type of risk which is not captured in the ICAPM specification. This test seems particularly relevant for this group of assets, given the efforts of the European Union toward integration of capital markets, also through harmonization of tax treatment for investors. The results are in table 3. A Wald test for their joint significance fails to reject the null hypothesis with a p-value of .83. In addition, a joint Wald test for the equality of the intercepts also

[18] On different datasets, Bollerslev, Engle and Wooldridge [1988] and Chan, Karolyi and Stulz [1992] estimate a GARCH-M with asset specific intercepts for a model which includes only market risk.

fails to reject the null (p-value = .75). Therefore the model under the null implies that there is no constant country specific risk or other type of friction that was left unaccounted in the specification.

Next, consider how risk is priced among European assets. The implication of the conditional version of the model is that optimizing investors update their strategy using the new available information in every period. Therefore, there is no reason to believe that the equilibrium reward per unit of risk in financial markets will stay constant in a conditional framework. In fact, the hypothesis of time-varying prices of risk can be tested, after adding specifications on how the parameters vary over time.

The prices of risk are assumed to be related to the information variables in a linear fashion:

$$\delta_{q,t-1} = \varphi_q' Z_{t-1} \quad q = 1, \ldots, L, m \tag{5.2}$$

where Z_{t-1} is a set of k information variables observed at the end of time $t-1$. Some of these instruments are common to all four prices of risk, while others are currency specific. The common instruments are the lagged dividend price ratio for the world index and the default premium. The price of market risk and the price of exchange risk of the Europound are also a function of the change in the US term premium and of the change in the US riskless rate.[19] For these two prices, I use US instruments since Harvey [1991] showed that they have power in predicting equity returns in foreign markets. On the other hand, the prices of exchange risk for the Eurofranc and the Eurolira also depend on the change in the German term premium and on the change in the German riskless rate.[20] This is because German instruments are likely to have predictive power on the prices of currencies that were members of the EMS. For example, while rejecting the hypothesis of strict German dominance, von Hagen and Fratianni [1990] uncover some asymmetries in the EMS since French and Italian monetary policies react more strongly to German policies than vice versa. Karfakis and Moschos [1990] find that German interest rates changes convey information about future movements in the interest rates of France and Italy.

The pricing of risk can be tested from $\phi_{q,t-1} = 0$ for all ϕs, while the time-variation in the prices of risk can be tested from $\phi_{q,t-1} = 0$ for all $\phi > 1$. In

[19] The previous variables are the commonly used instruments in the applied finance literature.

[20] All the instruments involving interest rates have not been lagged, since they are conditionally known at time $t-1$.

particular, if currency exposure does not represent a risk in financial markets, then the traditional version of the CAPM is sufficient for the pricing of securities. This hypothesis implies that $\phi_{l,t-1} = 0$ for $l = 1, \ldots, L$. To date, there exist only two examples of a conditional test for the international version of the CAPM which includes currency risk. With different methodologies, both Dumas and Solnik [1995] and De Santis and Gerard [1998] find support of a time-varying price of foreign exchange risk for the four largest world financial markets.

The results are in table 3. A Wald test for the significance of the time-varying parameters in the price of market risk rejects the null with a p-value of 0.002 and the joint test for the overall significance of the price also rejects the null hypothesis with a p-value of 0.0017. For currency risk, there is marginal evidence of time-variation in the prices (p-value 0.0764) but there is no conclusive evidence on the overall significance of currency risk (p-value 0.1336).[21]

The interesting evidence provided by the asset pricing model is that European markets have been effectively integrated. A common source of systematic risk, which is indication of integration in an asset-pricing framework, was significant even before the formal deadline set by the EU for the opening of legal barriers. The same sources of risk are thus explaining expected returns across different assets of the stock market and of the money market and there is no evidence of country specific risk. However these results are not surprising for a subset of countries whose efforts have been in developing very tight economic, political and financial linkages during the last two decades.

5.2. Structural Change in the Prices of Risk

In this paper the sample under investigation spans more than two decades: over such an extended period, the parameters are expected to be time-varying due to changing conditions in financial markets. In particular, by modifying the institutions governing financial markets, the liberalization process should also have caused a change in the assessment of the equilibrium reward to risk.

As one of the most important goals for the EU, this process of liberalization has been implemented through a series of gradual reforms directed at legally integrating European financial markets. These reforms were aimed at freeing capital

[21]A discussion of the individual coefficients is postponed to the last model specification. For reason of space I do not report all the individual estimates of the coefficients of most of the following specifications. They are available from the author upon request.

movements, liberalizing the ownership of financial assets, harmonizing accounting practices and market regulations, equalizing tax treatments and establishing transparency, free access and perfect competition. The implementation of the reforms has also increased the credibility of the commitment to liberalization and legal integration,[22] and should have consequently decreased political risk, that is risk of the imposition of future controls on the exchange of currencies and securities across borders.[23]

The ICAPM under investigation is now

$$r_{jt} = \sum_{l=1}^{X} \delta_{lt-1} cov_{t-1}[r_{jt}, r_{n+l,t}] + \delta_{mt-1} cov_{t-1}[r_{jt}, r_{mt}] + \epsilon_{jt} \tag{5.3}$$

which postulates time-varying prices for exchange and market risk based on the previous specification in (5.2). To the information set used in (5.2), I add a dummy which takes the value 1 after July 1990, the date set for the completion of liberalization for the members of the European Union. This structural change is introduced in all four prices of risk included in the model, since the event involved the elimination of both exchange and capital controls. This rather simplistic specification for a test on the relevance of liberalization would imply that an abrupt change on the pricing of risk happened on the day of the event, and the markets did not incorporate any expectation about the effects of liberalization prior to the deadline. Clearly, the credibility of such a scenario is subject to criticism and will be discussed later.[24]

In total, each price of risk is explained by five instruments, plus a constant.

Table 4 contains the results. Under the null hypothesis of constant prices of risk, all the ϕ parameters, except the constant, are set equal to zero. The robust Wald tests indicate that the null hypothesis can be rejected at the 5% significance level for the pricing of exchange risk, and at the .3% significance level for the

[22] As an example of this commitment, during the currency crisis of the fall 1992, those nations who had fully accomplished the liberalization of financial markets did not impose new border controls, even though they appeared to be an effective instrument against speculation.

[23] In the traditional literature on capital controls (see Aliber [1973], Claassen and Wyplosz [1982], and Obstfeld [1982]), the uncertainty on the imposition of future controls increases the risk for international diversification and generates an increase in the variance of rates of return.

[24] For example, Henry [2000] introduces a dummy variable for the eight-month window leading to the date of the liberalization. Bekaert and Harvey [2000] also use a liberalization indicator variable.

pricing of market risk. The prices of currency risk are also found to be jointly significant with a p-value of .02.

An alternative and more credible scenario assumes that markets had already taken into account the effects of legal integration by the opening date. This assumption is consistent with the gradual pace of the reforms for liberalization over the Eighties, aimed at progressively dismantling the barriers to free capital movements in France and Italy. Therefore, a smoother process rather than an abrupt effect is to be expected on the risk assessment of assets.[25]

The problem with this specification is in the choice of the instruments, since it is difficult to find variables that can capture the time-varying degree of intensity in capital barriers among financial markets. The offshore/onshore interest rate differential, or equivalently the covered interest differential, has provided in the past a useful approach to test whether a country has become integrated into international financial markets (see Dooley and Isard [1980], Claassen and Wyplosz [1982], Gultekin, Gultekin and Penati [1989]). Among those papers that focused on the European Monetary System, Rogoff [1985] identified capital controls as the cause of significant differentials between the Eurorate and the domestic rate for French and Italian assets. Giavazzi and Pagano [1988] used the covered interest rate differentials to investigate the effectiveness of capital controls for France and Italy.

In the data section, it was shown how the offshore/onshore differentials have been decreasing in the Eighties, in parallel with the progress of liberalization. Therefore, the one-month interest rate differential for each currency is used as information variable for each of the prices of exchange risk in substitution of the dummy. The price of market risk is parameterized with an equally weighted average of the three differentials, as an aggregate measure of the intensity of capital controls in Europe.[26] The other instruments are the same information variables of the previous specification. The results are in panel B of table 4. Under the null hypothesis of constant prices of risk, all the ϕ parameters, with the exception of the constant, are set equal to zero. The improvement of this specification is remarkable: the hypothesis of constant prices is rejected at any statistical level. Moreover, the robust Wald tests for the joint significance of the parameters reject

[25]For example, Bekaert and Harvey [2000] provide evidence that liberalizations are often gradual.

[26]Ferson and Harvey [1993, 1994] use various weighted averages as variables to represent global economic risk.

the null at any statistical level for all prices.

To verify the relevance of liberalization on the pricing of risk in European financial markets, a model with no exogenous information about the progress in liberalizing markets and no asset-specific intercepts was also estimated. Tests for the restrictions in this specification are in panel C of table 4. By comparing the p-values for specification of panel B and C, it can be inferred that the inclusion of information about liberalization increases the explanatory power of the instruments for the prices of risk. In fact, without the offshore/onshore differentials as instrumental variables, significance of the prices is somewhat reduced, although not eliminated.[27] Also, note that the overall significance of the prices of currency risk is increased from the specification in model (5.1), most likely due the removal of the intercepts from estimation.

Some of the instruments in the above specifications are not individually significant, which can be attributable to a degree of collinearity between variables or to high persistence of some instruments. For examples, the Ljung-Box statistics of panel A in table 2 revealed a high level of autocorrelation for the world index dividend price ratio and for the default premium. This implies that, in estimation, some of the instruments could not be adding any useful information.

For this reason, I estimate a new version of the model in (5.3) where the price of covariance risk is specified as

$$\delta_{m,t-1} = \exp(\varphi_m' Z_{t-1}) \tag{5.4}$$

In this case the set of information variable, Z_{t-1},contains a constant, the default premium, the change in the Eurodollar and the equally weighted average of the three differentials.[28] The price is thus a non-linear function of the instruments. This specification ensures that the price of risk will always be positive, as it is in the theoretical model, where the price of market risk depends on the relative wealth of each country and investors relative risk aversion.[29]

The prices of currency risk are still specified as

$$\delta_{l,t-1} = \varphi_l' Z_{t-1} \qquad l = 1, \ldots, L \tag{5.5}$$

[27] Differently from Dumas and Solnik [1995], De Santis and Gerard [1998] do not perform robustness checks to changes in the instruments for their methodology.

[28] These are the variables with the highest individual significance in my previous specifications.

[29] Most of previous literature specifies the price of covariance risk as linear. Only Bekaert and Harvey [1995], De Santis and Gerard [1997, 1998], Hardouvelis, Malliaropulos and Priestley [1999] use a non-linear specification.

Since the theoretical model does not preclude prices of currency risk from being negative, a linear specification is adopted. For the price of franc risk, Z_{t-1} includes a constant, the default premium, the change in the Euromark and the offshore/onshore one-month differential for France. For the price of lira risk, Z_{t-1} includes a constant, the default premium, the change in the Euromark and the offshore/onshore one-month differential for Italy. For the price of pound risk, Z_{t-1} includes a constant, the default premium, the change in the Eurodollar and the offshore/onshore one-month differential for the United Kingdom.

The estimates are reported in table 5. Panel A contains parameters of the return equation. Almost all the estimated parameters are significantly larger than their robust standard errors. Panel D has the results for tests of the model restrictions. Robust Wald tests for time variation in the reward for risk reject the null of a constant price at any significance level for all prices of currency risk and for the price of market risk. A joint test for significance of currency risk also rejects the null at any significance level. The results from the previous specifications are therefore remarkably robust to changes in the instruments and to changes in the specification of the price of market risk.[30]

In all the above specifications, I have parameterized second moments according to a GARCH (1,1) diagonal multivariate process. The raw data in table 1 show high level of autocorrelations in the series squared. The presence of GARCH effects is highly supported in all the estimated versions of the asset pricing model, and these effects are remarkably robust to all different specifications of the return equation of table 3, 4 and 5. I will thus comment only on the estimates of the model in table 5. All GARCH parameters are extremely significant: as previously documented in other work, the a_i are smaller than the b_i, an indication that lagged covariances have more weight than past innovations in explaining current covariances. Although most of the series show high persistence,[31] the condition for covariance stationarity[32] is satisfied in all cases.

The diagnostics on the normalized estimated residuals $z_t = \epsilon_t/\sqrt{h_t}$ are in panel B of table 5. The residuals still show evidence of non-normality in the skewness, kurtosis and Bera-Jarque statistics, indicating that the postulated GARCH process could not account for all the non-normality in the data. At the same time,

[30] To further check the robustness of my results, I estimated the specification in (6.4) and (6.5) with IFS data. The conclusion drawn from the estimated model did not change.

[31] Since $\alpha_i\alpha_j + \beta_i\beta_j$ is close to one (see Bollerslev and Engle [1986]).

[32] Since $\alpha_i\alpha_j + \beta_i\beta_j < 1 \forall i, j$ (see Bollerslev [1986]).

this supports the choice of quasi-maximum likelihood standard errors and of robust tests for inference. For all the estimated versions of the return equation, the Ljung-Box statistics fail to reject the null of no autocorrelation in the residuals for all assets. The same statistics performed on the residual squared fail to reject the null for all assets, except the German Index, as it is the case for all previous specifications and for the raw returns squared.

Additional diagnostics are provided by the pricing errors and by the root mean squared errors, as measure of the difference between actual and predicted returns, conditional on a correct specification for the model. In general, the average pricing errors are a fraction of the observed average returns. However, the estimated model generates a large error for the Italian index, an indication that this country performed worse than expected, given its level of riskiness.

In conclusion, the evidence presented in this section strongly suggests that liberalization has affected the dynamics of the prices of exchange and market risk. To analyze the resulting patterns, I now turn to a description of the estimated variables.

5.3. Liberalization and time-varying risk premia

Panel A of table 6 contains averages of the estimated risk premia for the specification of (5.3) with (5.4) and (5.5). The series have been split into subsamples at July 1, 1990, the formal deadline set by the European Union for the completion of markets liberalization. However, since the operation of the EMS represents another potential structural change in the dynamics of risk, I also divide the series at the inception of the EMS on March 1979.

Insights can be offered by a comparison of the relative magnitude of exchange risk premia over the traditional premia for market risk. De Santis and Gerard [1998] establish that in the four largest world capital markets, the contribution of the currency risk component in total risk premia is undetected when average premia are considered, but is instead relevant in a conditional framework. As expected, market premia are larger for stocks, while exchange premia are relatively bigger for currency deposits. Therefore, an international asset pricing model without exchange rate as a source of risk would be mispecified in particular for the pricing of assets of the money market. The size of currency premia is economically significant, a conclusion similar to the one offered in De Santis and Gerard [1998]. However, in that paper, the size of estimated currency premia is generally larger, which is not surprising given that currencies in their investigation are free floating. [33] Time-varying prices and risk premia are plotted in figures 2 and 3. While in table 6 exchange risk premia can appear to be quite small on average due to fluctuating negative and positive values over long samples, the plots reveal that exchange risk premia are instead sizable over periods.

Now consider the impact of changing institutions on the assessment of risk in European financial markets.

First, consider market risk premia. In panel B, the mean and standard deviation of the time-varying price of market risk, $\delta_{m,t-1}$, are found to be lower after the completion of liberalization. The average of the price decreases from 3.86 to 1.5: the reward-to-risk has thus decreased in European financial markets. Risk sharing through increased diversification opportunities can explain this fall in the expected compensation for risk (see Stulz [1999]).

In panel A, the estimated market premia vary considerably and are found to be consistently lower across all assets in the second half of the subsample. Since the estimated model is fully parameterized, I can separate the impact of the dynamics of the price of risk from the estimated covariances. This analysis reveals that the covariances are basically unchanged after both events. The fall in the price of market risk dominates the dynamics of market risk premia.[34] On the other hand, a uniform pattern cannot be inferred from the EMS subsamples. If anything, risk premia seem larger, most likely as a consequence of a higher price of risk over this subsample. These conclusions are confirmed by inspection of the plots in figure 2 and 3.

Now consider exchange risk premia. Panel B provides some statistics on the estimated time-varying prices of currency risk. Since the theory does not impose restrictions on the sign of these prices, they are a linear function of the information variables and fluctuate between positive and negative values. Therefore, it is more difficult to draw conclusions only from such statistics and one needs to rely on the inspection of their path in figure 2 to gain further insights. All prices are

[33] The findings of this paper are in contrast with the results in Ayuso and Restoy [1996]. In their paper, for example, the Eurofranc premium is estimated in the range of 28 basis point (annualized) at its highest among the subperiods, as compared to my estimate of 170 basis points over a comparable period. A possible explanation for the discrepancy is due to the methodology adopted for estimation since I use a GARCH specification to model conditional second moments.

[34] For reason of space, I do not report separate estimates for the time-varying covariances but they are available from the author.

smaller in absolute value and have lower standard deviation after the completion of liberalization. On the contrary, no uniform direction can be found for the prices after the inception of the EMS. First, the plots confirm the evidence of a decreasing trend for almost all prices. The volatility of prices is also clearly reduced in the last decade. Second, prices of currency risk often tend to assume large negative values in correspondence of large values of the instruments. For example, compare the pattern of the French offshore/onshore differential in figure 1 with the dynamics of the price of Eurofranc risk. The differential widened in correspondence with the increase in capital controls after 1981, with the turbulence among currencies which lead to the devaluation of April 86, and with the currency crisis of August 93. These same events caused large negative values for the price of Franc risk and consequently, with positive covariances, bigger negative exchange risk premia. Prices of risk are thus sensitive to uncertainty in financial markets which could lead to slowing or reversal of the liberalization process, as represented by tightening of capital controls or by periods of strain in the EMS. This also implies that, with negative currency premia, international investors are willing to give up some of the total risk premium over these periods when the hedging value of the assets in the portfolio becomes predominant.

Turning to the estimated premia, while no consistent pattern can be inferred after the liberalization was completed, exchange risk premia are consistently lower after the inception of the EMS. Therefore, there is evidence of lower exchange risk premia, mostly due to the reduction in volatility obtained by keeping currencies movements within predetermined boundaries. Lower covariances have thus implied consistently lower currency premia in the EMS subsample. The reduction is really dramatic for assets in the money market after 1979 (start of the EMS), but there is also evidence of it in the equity market, as illustrated by averages and plots of the assets estimated premia.

In panel A, the size of total risk premia offers the overall picture. For assets in the stock market, total risk premia are shown to be consistently lower in the period following the liberalization deadline and basically unchanged after the EMS inception. Assets in the money market are shown to be more affected by the operation of the EMS which has caused a general reduction in total premia, but they also provide evidence of a decrease after the liberalization.[35] The conclusions

[35] The significant decrease in the total premia for the UK equity in the EMS subsample can be easily explained since the elimination of border restrictions in UK in July 1979 almost coincides with the inception of the EMS in March 1979.

drawn from the statistics of the estimated total risk premia are confirmed after comparing the plots of the assets premia in figure 3.

As a final check, I use the estimated total risk premia in a cross-sectional time-series regression model using as regressor a dummy variable which takes the value of 1 at the official liberalization date. I estimate the model for three groups, one which includes all assets, a second one with only assets from the stock market and a third one with only assets from the money market. The pooled regressions allow for fixed effects and for heteroskedasticity and autocorrelation in the residuals. Table 7 reports the results. As expected, all the estimated slope coefficients are negative and they are much larger than the corresponding standard errors. The results imply that on an annualized basis the expected risk premia have decreased by an average of 4% across all markets. This simple test corroborates the conclusion that a structural change in the risk premia is related to the liberalization of European capital markets.

In summary, during the completion of liberalization, an overall decrease in the expected compensation for risk has caused a general reduction in total risk premia. The decrease in market risk premia is generally a result of a lower price of risk while the decrease in currency premia is also driven by smaller covariances. Overall, a decreasing trend in total risk premia is confirmed across all assets. Equity and currency markets are thus now less risky for European investors than they were two decades ago.

6. Conclusions

This paper uses the International CAPM of Adler and Dumas [1983] to analyze the impact of liberalization on financial markets of the European Union. Estimation is performed using the approach proposed by De Santis and Gerard [1997, 1998]. The analysis provides evidence that the model can help in assessing the relevant sources of risk for European assets and in determining the changes in the expected compensation for risk.

The results can be summarized as follows. European financial markets are found to be effectively integrated over the whole sample from March 1974 to August 1995, even before the openings of legal barriers set by the European Union. Only one price of market risk exists and the same sources of risk consistently price all assets in the equity market as well as in the Eurocurrency market. The result could be attributable to the strong economic links from trading developed by these countries over the last two decades. The analysis also provides evidence that currency risk is priced and the size of currency premia is economically significant. Hence, European investors are rewarded for their exposure to exchange risk.

Prices of risk are time-varying. The evolution of risk prices and risk premia throughout the institutional changes is rather interesting. Time-variation in the expected compensation for risk is consistent with the major financial markets reforms of the last two decades. Liberalization is characterized by a decreasing trend in both exchange and market risk premia mainly due to a gradual decline in the prices of risk during the liberalization process. Risk sharing among economic agents may provide an explanation for this finding. There is evidence that markets tend to react to periods of uncertainty in the progress toward the completion of liberalization. Tightening of capital controls or times of strain in the EMS are examples of these occurrences. However, there is also evidence that in such periods international assets provide a good hedge to uncertainty. The operation of the European Monetary System has also had a significant impact, by decreasing exchange risk premia over time. As a consequence, equity and currency markets are less risky overall for European investors than they were two decades ago.

The validity of the pricing restrictions of the International CAPM has also important policy implications for the debate on the single currency in Europe. One of the arguments in favor of the introduction of a single currency is the potential to reduce risk in financial markets.[36] In fact, currency risk is sometimes considered an obstacle to international diversification. However, I found evidence that investors are indeed compensated for their exposure to currency risk. This implies that the elimination of national currencies will only lead to a reduction in the number of assets available for international diversification. Thus, the case for European Monetary Union has to be based on grounds other than the reduction of risk.

[36] See for example, the opinion of the European Commission [1992].

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Table 1: Summary Statistics for Asset Returns

PANEL A: Distributional Statistics

Statistics for asset returns. Equity indices are from MSCI while eurorates are from DRI. Returns are monthly percentage, denominated in DM and in excess of the Euro-DM one month deposit rate. The period is from March 1974 to August 1995 with 258 observations. The test for the kurtosis coefficient has been normalized to zero, B-J is the Bera-Jarque test for normality based on excess skewness and kurtosis, Q is here the Ljung-Box test for autocorrelation of order 12 for the returns and for the returns squared.

	France	Germany	Italy	U.K.	Euro FF	Euro LR	Euro BP	World
weights[a]	0.034	0.042	0.015	0.108				1.000
Mean	0.462	0.432	0.155	0.668	0.146	0.133	0.06	0.443
Std. Dev.	6.56	5.15	8.11	7.56	1.23	2.20	2.71	4.06
Skewness	0.06	-0.46**	0.45**	.93**	-.73**	-1.38**	-0.11	-.54**
Kurtosis	1.42**	2.54**	1.05**	8.44**	4.09**	8.71**	1.39**	2.38**
B-J	20.4**	75.08**	19.5**	769.4**	194**	862**	20.04**	70.29**
$Q(z)_{12}$	10.18	14.31	13.37	20.35	31.84**	10.17	16.42	13.74
$Q(z^2)_{12}$	20.46*	42.38**	17.01	16.11	87.24**	33.12**	7.42	4.35

* significant at the 5% level. ** significant at the 1% level.

[a] as of December 31,1990

PANEL B: Crosscorrelations of squared returns between world and asset shown in the column

Lag	France	Germany	Italy	U.K.	Euro FF	Euro LR	Euro BP
-6	-.022	.025	-.029	-.028	.020	-.025	-.018
-5	-.035	.096	-.022	-.018	.011	.021	-.104
-4	0.142*	.004	-.073	-.041	.031	-.064	-.074
-3	-.007	.069	-.007	.007	.003	-.066	-.058
-2	-.058	-.056	-.022	-.035	.055	-.054	-.039
-1	.082	.056	0.154*	.011	.026	-.032	-.059
0	0.495**	0.585**	0.326**	0.417**	.037	.007	.083
1	-.004	0.22*	-.091	.021	.040	-.040	-.055
2	.035	-.023	-.050	.036	.003	-.016	-.033
3	.051	.048	.012	.011	.002	.005	-.011
4	0.244*	0.228*	-.015	.105	-.006	-.041	.000
5	.036	.066	-.021	.069	-.045	-.077	-.005
6	.072	.000	-.050	.016	-.035	-.016	-.036

* significant at the 5% level. ** significant at the 1% level.

PANEL C: Pairwise Correlation for Asset Returns

	France	Germany	Italy	U.K.	Euro FF	Euro LR	Euro BP	World
France	1	0.496	0.404	0.518	0.301	0.108	0.104	0.576
Germany		1	0.344	0.375	-0.019	0.08	0.04	0.456
Italy			1	0.353	0.096	0.445	0.254	0.485
U.K.				1	0.253	0.095	0.45	0.682
Euro FF					1	0.391	0.299	0.208
Euro LR						1	0.393	0.205
Euro BP							1	0.324
World								1

Table 2: Summary Statistics for Information Variables

PANEL A: Distributional Statistics

Statistics for information variables. Their symbols are as follows:
XDPR is the lagged world index dividend yield in excess of the Euro-DM one month deposit rate.
ΔUSTP is the change in the US term premium measured by the yield difference between the 10-year Treasury Notes and the 3-month Treasury Bills.
ΔGETP is the change in the German term premium mesured by the yield difference between the Government Bonds with an average remaining life to maturity of more than three years and the 12-month Treasury Bills of the Federal debt
DP is the US default premium measured by the yield difference between Moody's Baa and Aaa rated bonds
ΔUS$ is the change in the one-month Euro$ deposit rate
ΔDM is the change in the one-month EuroDM deposit rate
FDF is the offshore/onshore one-month interest rate differential for France, in percentage
IDF is the offshore/onshore one-month interest rate differential for Italy, in percentage
UKDF is the offshore/onshore one-month interest rate differential for the United Kingdom, in percentage
The period is from March 1974 to August 1995 with 258 observations.

	XDPR	ΔUSTP	ΔGETP	DP	ΔUS$	ΔDM	FDF	IDF	UKDF
mean	-0.250	0.007	0.004	1.220	-0.0006	-0.001	0.112	0.115	0.101
St.dev.	0.217	0.496	0.637	0.468	0.109	0.078	0.240	0.562	0.125
$Q(z)_{12}$	1989**	58.4**	59.87**	1636**	89.59**	138**	167.6**	82.11**	131.5**

PANEL B: Autocorrelations

	XDPR	ΔUSTP	ΔGETP	DP	ΔUS$	ΔDM	FDF	IDF	UKDF
ρ_1	0.926	0.246	-0.389	0.952	-0.211	-0.387	0.388	0.135	0.344
ρ_2	0.913	-0.170	0.001	0.888	-0.054	-0.130	0.333	0.178	0.267
ρ_3	0.913	-0.048	-0.022	0.839	0.220	0.375	0.202	0.165	0.307
ρ_4	0.865	-0.087	0.021	0.803	-0.350	-0.203	0.115	0.009	0.265
ρ_5	0.849	-0.088	-0.002	0.768	0.077	0.092	0.109	0.099	0.190
ρ_6	0.812	-0.145	0.112	0.714	-0.012	0.081	0.219	0.030	0.180
ρ_{12}	0.604	-0.190	0.047	0.494	0.072	0.314	0.054	0.087	0.060

PANEL C: Pairwise Correlation for Instrumental Variables

	XDPR	ΔUSTP	ΔGETP	DP	ΔUS$	ΔDM	FDF	IDF	UKDF	AvDF
XDPR	1	-0.132	-0.017	0.03	0.151	0.178	-0.021	0.0482	0.026	0.037
ΔUSTP		1	-0.031	0.141	-0.362	-0.069	0.042	-0.079	-0.006	-0.052
ΔGETP			1	0.008	-0.053	-0.071	0.049	0.23	0.032	0.217
DP				1	-0.157	-0.071	0.258	0.232	0.0346	0.294
ΔUS$					1	0.548	0.089	0.082	0.215	0.141
ΔDM						1	0.0459	0.0462	0.178	0.088
FDF							1	0.108	0.276	0.502
IDF								1	0.095	0.895
UKDF									1	0.367
AvDF										1

TABLE 3: Specification Tests for Pricing Restrictions

Tests of the following restrictions in the mean equation of the ICAPM model:

$$R_{j,t} = \alpha_i + \delta_{m,t-1}\, cov\, (R_{jt}, R_{mt} \mid \vartheta_{t-1}) + \Sigma^{3}_{l=1}\, \delta_{l,t-1}\, cov\, (R_{jt}, R_{lt} \mid \vartheta_{t-1}) + \varepsilon_{jt}$$

where $R_{j,t}$ are asset excess returns, δ_m is the price of world covariance risk, δ_l are the prices of currency risk and $e_t \mid J_{t-1} \sim N(0, Ht)$. The time-varying prices are given by:

$$\delta_{q,t-1} = \phi_q' Z_{t-1} \qquad q = m, FF, IL, BP$$

where Z is a set of k information variables which vary with each specification. A comprehensive list of these instruments includes the world index dividend yield in excess of the risk-free rate (XDPR), the change in the term premium for the US (ΔUSTP) and for Germany (ΔGETP), the default premium (DP), the change in the Euro$ (ΔUS$), the change in the EuroDM (ΔDM).
Only the robust Wald statistics and their P-values are reported with the corresponding restrictions.
In all cases, the time-varying conditional covariance H_t is parameterized as:

$$H_t = H_0 * (\iota\iota' - \mathbf{aa}' - \mathbf{bb}') + \mathbf{aa}' * \Sigma_{t-1} + \mathbf{bb}' * H_{t-1},$$

where * denotes the Hadamard product, **a** and **b** are (N x 1) vector of constants, ι is (N x 1) unit vector, and Σ_{t-1} is the matrix of cross error terms, $\varepsilon_{t-1}\varepsilon'_{t-1}$.
The equity data are from MSCI, while the deposits rates are from DRI. All returns are denominated in DM.

Null Hypothesis	χ^2	d.f.	p-value
$\alpha_j = 0$, for all j	4.24	8	0.8344
$\alpha_j = \alpha$, for all j	4.17	7	0.7595
$\phi_{l,k} = 0$, for all l, k	21.1	15	0.1336
$\phi_{l,k} = 0$, for all l, k>1	19.53	12	0.0764
$\phi_{m,k} = 0$, for all k	19.26	5	0.0017
$\phi_{m,k} = 0$, for k>1	16.51	4	0.0024

TABLE 4: Tests on the Relevance of Liberalization

Results from robust Wald tests concerning the relevance of liberalization. The estimated model is:

$$R_{j,t} = \delta_{m,t-1}\, cov\,(R_{jt},R_{mt} \mid \vartheta_{t-1}) + \Sigma^{3}_{l=1}\, \delta_{l,t-1}\, cov\,(R_{jt},R_{lt} \mid \vartheta_{t-1}) + \varepsilon_{lt}$$

where $R_{j,t}$ are asset excess returns, δ_m is the price of world covariance risk, δ_l are the prices of currency risk and $\varepsilon_t \mid \vartheta_{t-1} \sim N(0, H_t)$. The time-varying prices are given by:

$$\delta_{q,t-1} = \phi_q' \mathbf{Z}_{t-1} \qquad q = m, FF, IL, BP$$

where **Z** is a set of *k* information variables which vary with each specification. A comprehensive list of these instruments includes the world index dividend yield in excess of the risk-free rate (XDPR), the change in the term premium for the US (ΔUSTP) and for Germany (ΔGETP), the default premium (DP), the change in the Euro$ (ΔUS$), the change in the EuroDM (ΔDM), a dummy variable which takes the value 1 after July 1, 1990 (Dummy), the offshore/onshore differentials for France (FDF), Italy (IDF) and the United Kingdom (UKDF) and an average offshore/onshore differential (AvDF). The set of information variables used with a particular specification is reported under each price.

The time-varying conditional covariance H_t is parameterized as:

$$H_t = H_0 * (\iota\iota' - \mathbf{aa}' - \mathbf{bb}') + \mathbf{aa}' * \Sigma_{t-1} + \mathbf{bb}' * H_{t-1},$$

where * denotes the Hadamard product, **a** and **b** are (N x 1) vector of constants, ι is (N x 1) unit vector, and Σ_{t-1} is the matrix of cross error terms, $\varepsilon_{t-1}\varepsilon'_{t-1}$.

The models are estimated by Quasi-Maximum Likelihood.

δ_m	δ_{FF}	δ_{IL}	δ_{BP}	Null Hypothesis	χ^2	d.f.	p-value
				Panel A			
XDPR	XDPR	XDPR	XDPR				
ΔUSTP	ΔGETP	ΔGETP	ΔUSTP	$\phi_{l,k} = 0$, for all l, k	31.24	18	0.0269
DP	DP	DP	DP	$\phi_{l,k} = 0$, for all l, k>1	24.96	15	0.0505
ΔUS$	ΔDM	ΔDM	ΔUS$	$\phi_{m,k} = 0$, for k>1	17.92	5	0.0030
Dummy	Dummy	Dummy	Dummy				
				Panel B			
XDPR	XDPR	XDPR	XDPR				
ΔUSTP	ΔGETP	ΔGETP	ΔUSTP	$\phi_{l,k} = 0$, for all l, k	53.82	18	0.0000
DP	DP	DP	DP	$\phi_{l,k} = 0$, for all l, k>1	47.11	15	0.0000
ΔUS$	ΔDM	ΔDM	ΔUS$	$\phi_{m,k} = 0$, for k>1	21.6	5	0.0006
AvDF	FDF	IDF	UKDF				
				Panel C			
XDPR	XDPR	XDPR	XDPR				
ΔUSTP	ΔGETP	ΔGETP	ΔUSTP	$\phi_{l,k} = 0$, for all l, k	24.92	15	0.0509
DP	DP	DP	DP	$\phi_{l,k} = 0$, for all l, k>1	20.15	12	0.0642
ΔUS$	ΔDM	ΔDM	ΔUS$	$\phi_{m,k} = 0$, for k>1	17.59	4	0.0015

TABLE 5: ICAPM with Time-varying Prices of Risk

Panel A: the Mean Equation

The estimated model is:

$$R_{j,t} = \delta_{m,t-1}\, cov\,(R_{jt},R_{mt} \mid \vartheta_{t-1}) + \Sigma^{3}_{l=1}\, \delta_{l,t-1}\, cov\,(R_{jt},R_{lt} \mid \vartheta_{t-1}) + \varepsilon_{jt}$$

where $R_{j,t}$ are the assets excess returns, δ_m is the price of world covariance risk, δ_l are the prices of currency risk and $\varepsilon_t \mid \vartheta_{t-1} \sim N(0, H_t)$. The time-varying prices are estimated each with a different set of conditioning information. Price specifications are given by:

$$\delta_{m,t-1} = exp\,(\,\phi_m' \mathbf{Z}_{t-1}\,)$$

Z is a set of *k* information variables which includes a constant, the default premium (DP), the change in the Euro$ (ΔUS$) and an average offshore/onshore differential (AvDF) ,

$$\delta_{FF,t-1} = \phi_{FF}' \mathbf{X}_{t-1}$$

X is a set which includes a constant, the default premium (DP), the change in the EuroDM (ΔDM) and the offshore/onshore French differential (FDF),

$$\delta_{IL,t-1} = \phi_{IL}' \mathbf{Y}_{t-1}$$

the set **Y** includes a constant, the default premium (DP), the change in the EuroDM (ΔDM) and the offshore/onshore Italian differential (IDF),

$$\delta_{BP,t-1} = \phi_{BP}' \mathbf{K}_{t-1}$$

K is a set which includes a constant, the default premium (DP), the change in the Euro$ (ΔUS$) and the offshore/onshore differential for the United Kingdom (UKDF).

The equity data are from MSCI, while the deposits rates are from DRI. All returns are denominated in DM. There are 258 observations, from March 1974 to August 1995.

The model is estimated by Quasi-Maximum Likelihood: heteroskedasticity-consistent standard errors are reported in parentheses.

Log Likelihood: -3661.56

	δ_m	δ_{FF}	δ_{IL}	δ_{BP}
constant	-5.5837	0.3735	-0.2048	0.0016
	(1.0293)	(.1581)	(.1151)	(.0799)
DP	1.5928	-0.1591	0.1975	-0.0466
	(.4288)	(.1189)	(.0822)	(.0673)
ΔUS$	-1.9407			0.5766
	(.8591)			(.2615)
ΔDM		0.9776	-0.6691	
		(.5166)	(.4238)	
AvDF	-1.3471			
	(1.2615)			
FDF		-0.6543		
		(.1831)		
IDF			-0.1225	
			(.0716)	
UKDF				0.297
				(.1471)

Panel B: Covariance Equation

H_t is the time-varying conditional covariance parameterized as:

$$H_t = H_0 * (\iota\iota' - \mathbf{aa}' - \mathbf{bb}') + \mathbf{aa}' * \Sigma_{t-1} + \mathbf{bb}' * H_{t-1},$$

where * denotes the Hadamard product, **a** and **b** are (N x 1) vector of constants, ι is (N x 1) unit vector, and Σ_{t-1} is the matrix of cross error terms, $\varepsilon_{t-1}\varepsilon'_{t-1}$. Heteroskedastic consistent standard errors are reported in parentheses.

	France	Germany	Italy	U.K.	Euro FF	Euro LR	Euro BP	World
a_i	0.1077	0.1381	0.1316	0.2426	0.5062	0.3274	0.1550	0.2758
	(.0395)	(.0495)	(.0462)	(.0395)	(.0480)	(.0653)	(.0809)	(.0463)
b_i	0.9878	0.9761	0.9543	0.9469	0.7755	0.8158	0.8748	0.7704
	(.0154)	(.0072)	(.0339)	(.0222)	(.0533)	(.0394)	(.0659)	(.0869)

Panel C: Model Diagnostics for standardized residuals

The test for the kurtosis coefficient has been normalized to zero, B-J is the Bera-Jarque test for normality based on excess skewness and kurtosis, Q is here the Ljung-Box test for autocorrelation of order 12 for the residuals and the residuals squared.

	France	Germany	Italy	U.K.	Euro FF	Euro LR	Euro BP	World
Average Excess Ret.	0.462	0.432	0.155	0.668	0.146	0.133	0.06	0.443
Average Pricing Error	-0.335	0.083	-0.524	-0.190	0.017	-0.023	-0.041	-0.404
RMSE	6.61	5.14	8.05	7.48	1.25	2.21	2.66	4.53
Skewness	0.04	-.35*	.51**	0.20	-0.92**	-2.49**	-0.14	-.60**
Kurtosis	1.55**	1.73**	1.21**	4.21**	2.25**	14.33**	1.41**	3.39**
B-J	24.22**	35.36**	25.9**	183.6**	87.75**	2381**	20.69**	132.7**
$Q_{12}(z)$	9.26	13.22	14.97	12.34	18.20	9.43	17.23	12.31
$Q_{12}(z^2)$	19.24	26.98**	17.59	8.88	11.60	1.42	4.41	2.57

* significant at the 5% level. ** significant at the 1% level.

Panel D: Specification Tests

Robust Wald tests	Null Hypothesis	χ^2	d.f.	p-value
for exchange rate risk	$\phi_{c,k} = 0$, for all c, k	72.94	12	0.0000
for time-varying exchange risk	$\phi_{c,k} = 0$, for all c, k>1	41.76	9	0.0000
for time-varying market risk	$\phi_{m,k} = 0$, for k>1	27.86	3	0.0000

Table 6: Comparison of Estimated Risk Premia

Averages of the estimated risk premia and prices of risk for the following model specification:

$$R_{j,t} = \delta_{m,t-1}\, cov\ (R_{jt}, R_{mt} \mid \vartheta_{t-1}) + \Sigma^{3}_{l=1}\, \delta_{l,t-1}\, cov\ (R_{jt}, R_{lt} \mid \vartheta_{t-1}) + \varepsilon_{jt}$$

with time-varying prices of risk, specified as in table 5.
Each panel reports market risk premia, exchange risk premia, total premia and prices of risk over four subperiods: before and after the deadline of July 1, 1990 to complete the liberalization process for European financial markets, before and after March 1979, the inception of the European Monetary System.

Panel A

	France	Germany	Italy	U.K.	Euro FF	Euro LR	Euro BP	World
				Market Risk Premia				
pre-liber.	0.663	0.392	0.678	0.9	0.034	0.06	0.156	0.78
post-liber.	0.269	0.181	0.285	0.341	0.01	0.031	0.059	0.32
pre-EMS	0.502	0.282	0.47	0.774	0.032	0.032	0.104	0.561
EMS	0.588	0.36	0.619	0.762	0.027	0.06	0.141	0.703
				Exchange Risk Premia				
pre-liber.	0.176	0.027	0.209	0.073	0.088	0.167	-0.029	0.081
post-liber.	0.402	-0.055	-0.267	0.152	0.139	-0.105	-0.038	0.022
pre-EMS	0.293	0.022	0.152	0.447	0.194	0.14	0.152	0.196
EMS	0.208	0.002	0.077	-0.017	0.071	0.09	-0.088	0.026
				Total Risk Premia				
pre-liber.	0.836	0.419	0.887	0.973	0.123	0.228	0.127	0.862
post-liber.	0.671	0.125	0.019	0.493	0.150	-0.074	0.020	0.342
pre-EMS	0.796	0.305	0.623	1.223	0.227	0.173	0.257	0.758
EMS	0.797	0.362	0.696	0.745	0.099	0.150	0.053	0.730

Panel B

	Prices of risk			
		δ_m		
	mean	st.dev.	min.	max.
pre-liberalization	3.86	4.41	0.18	32.01
post-liberalization	1.5	0.74	0.69	4.73
pre-EMS	2.72	2.76	0.18	14.28
EMS	3.47	4.26	0.69	32.01
		δ_{FF}		
	mean	st.dev.	min.	max.
pre-liberalization	6.48	21.66	-91.04	57.91
post-liberalization	22.86	8.04	-20.45	41.99
pre-EMS	5.98	15.75	-55.12	27.4
EMS	11.79	21.61	-91.04	57.91
		δ_{IL}		
	mean	st.dev.	min.	max.
pre-liberalization	4.11	12.71	-41.82	76.25
post-liberalization	-3.36	5.91	-15.88	14.38
pre-EMS	0.52	16.42	-41.82	76.25
EMS	2.87	10.18	-22.17	39.97
		δ_{BP}		
	mean	stdev.	min.	max.
pre-liberalization	-2.77	9.48	-38.5	30.58
post-liberalization	-1.81	4.29	-11.4	10.45
pre-EMS	0.73	10.18	-18.66	28.24
EMS	-3.55	7.7	-38.5	30.58

TABLE 7: Structural Changes in Estimated Risk Premia

Estimates of the following cross-sectional time-series regression model:

$$R^{e}_{j,t} = \beta_0 + \beta_1 D_{t+} \varepsilon_{j,t}$$

where $R^{e}_{j,t}$ are the estimated asset total risk premia from the model in table 5 and D_t is a dummy variable which takes the value 1 after July 1, 1990.
Heteroskedastic and autocorrelation robust standard errors are reported under each coefficient. The pooled regressions have been estimated with fixed effects. The estimates are in percent per month.

Group	β_1
All assets	-0.3387
	0.0531
Equity indices only	-0.4520
	0.0874
Eurocurrency deposits only	-0.1274
	0.0510

FIGURE 1: Differentials

Figure 1a: French Offshore/Onshore Differential

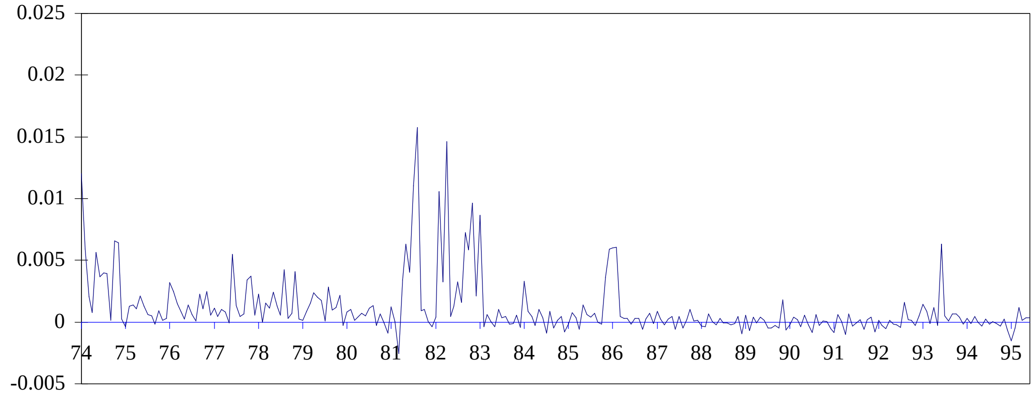


Figure 1b: Italian Offshore/Onshore Differential

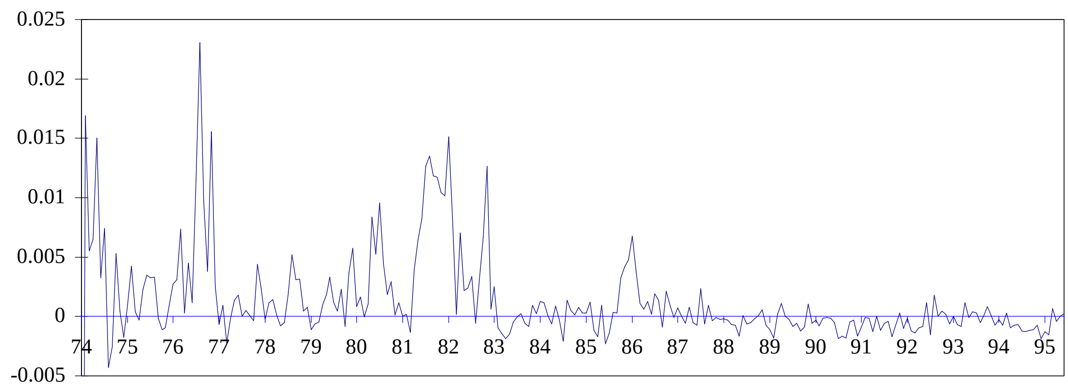


Figure 1c: U.K. Offshore/Onshore Differential

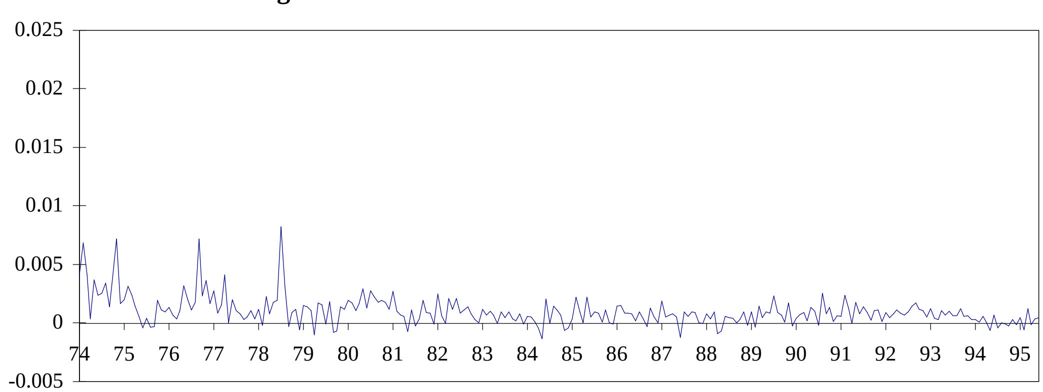
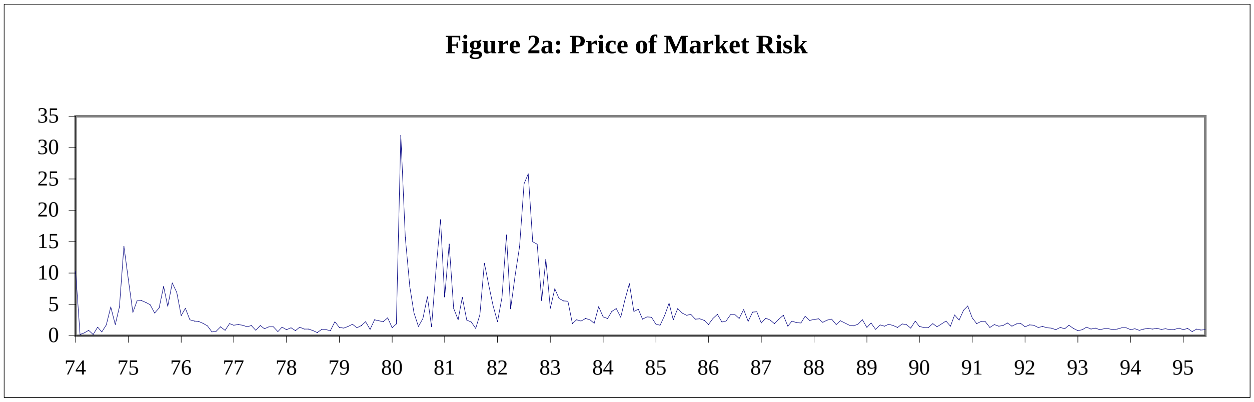
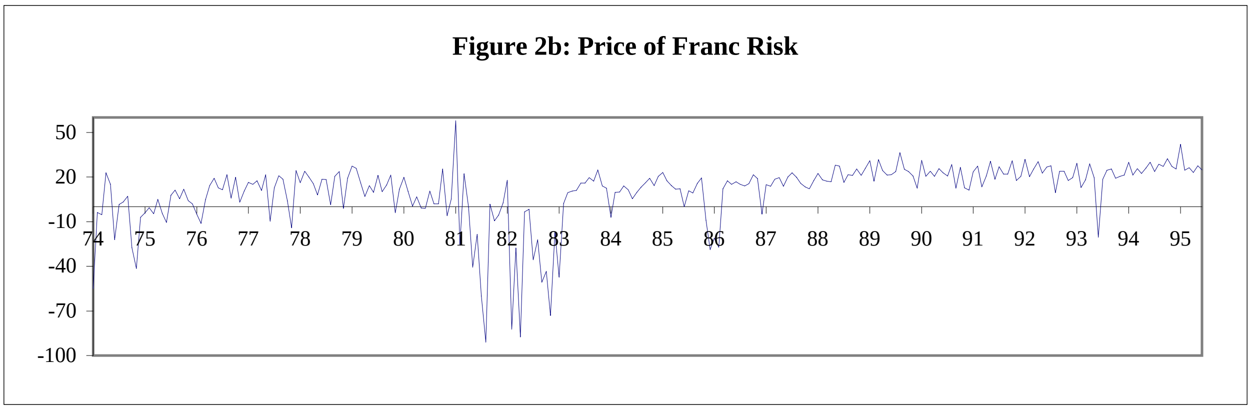
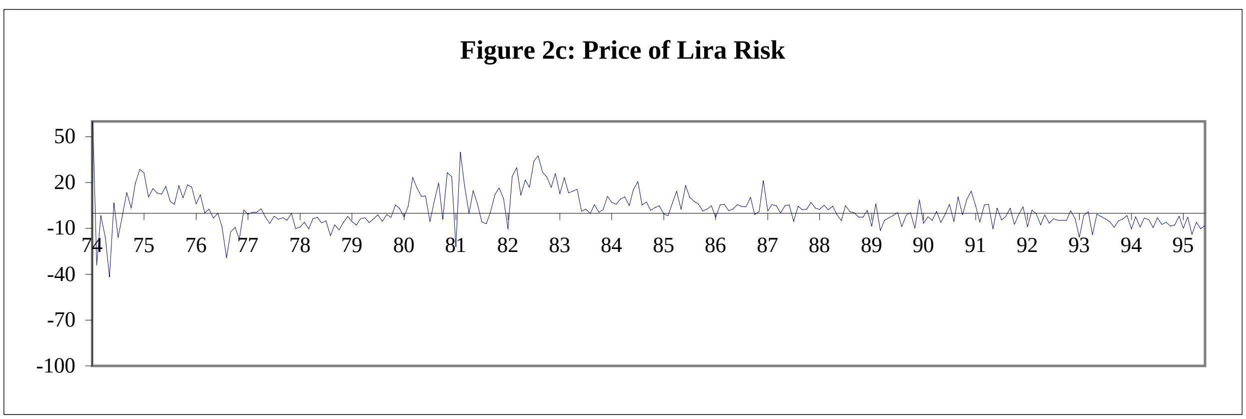


FIGURE 2: Prices of Risk







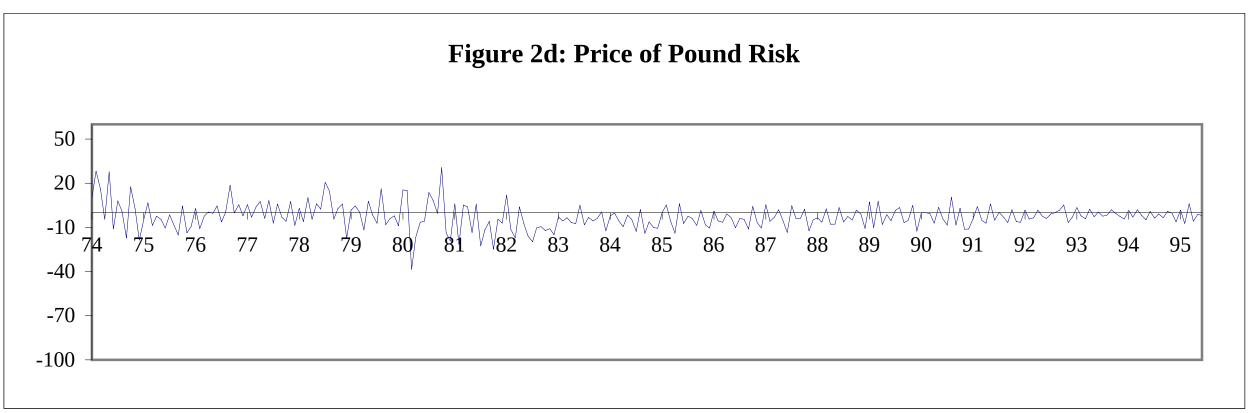
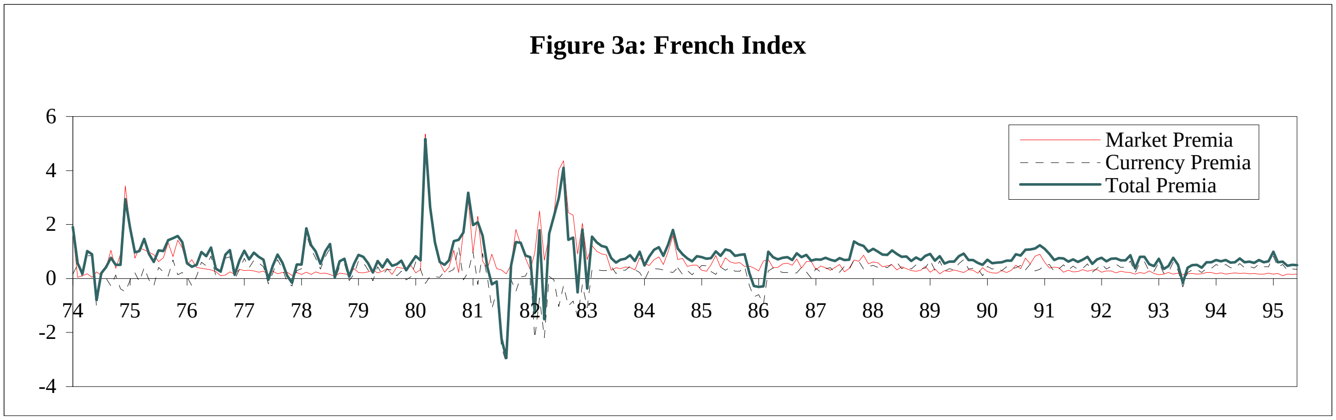
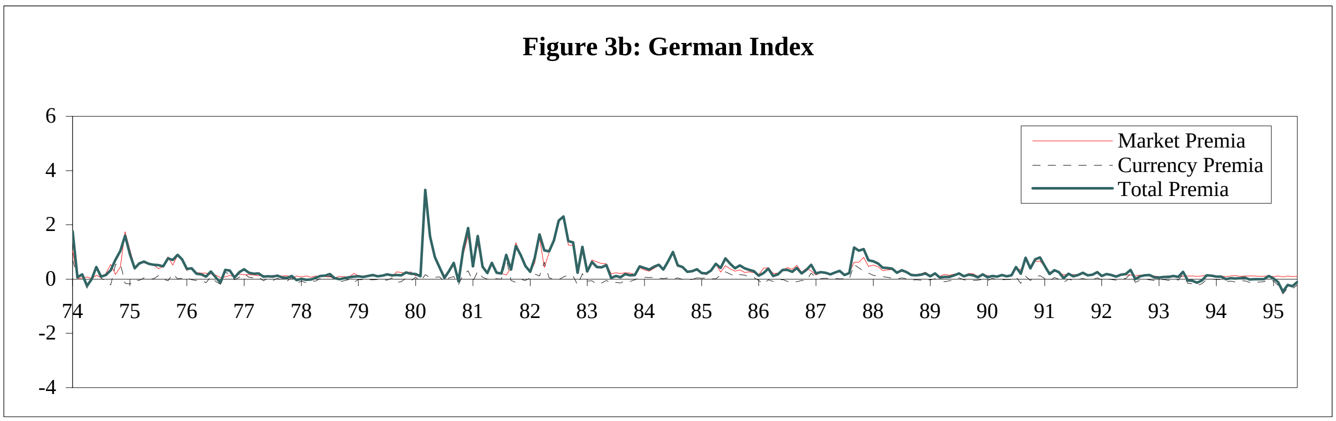
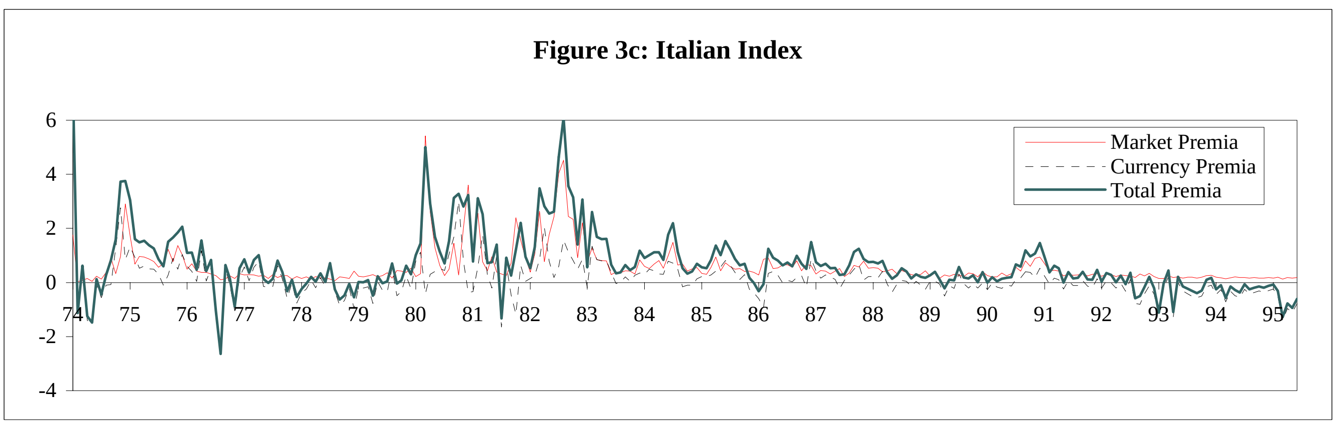


FIGURE 3: Estimated Risk Premia







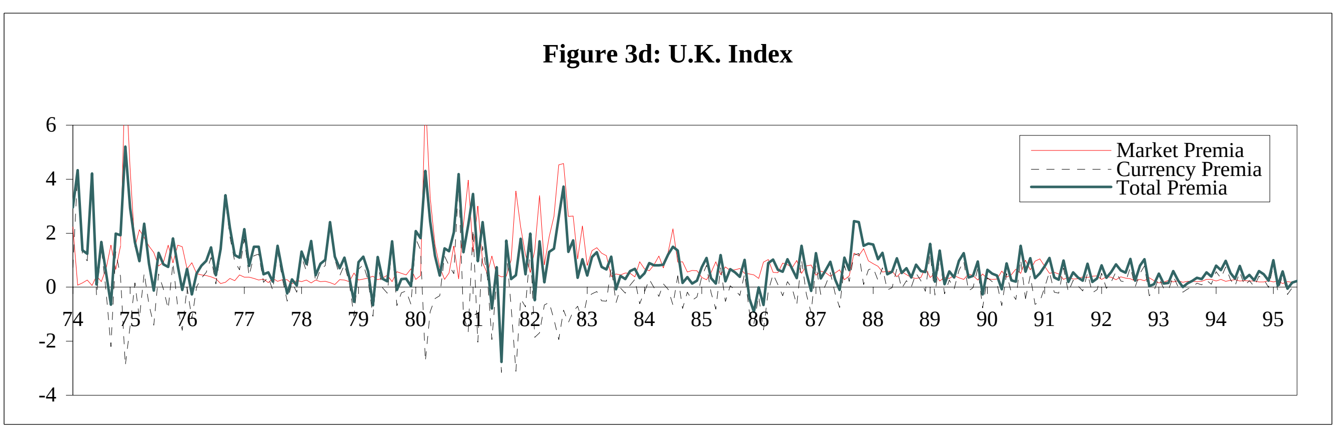


FIGURE 3: Estimated Risk Premia

Figure 3e: World Index

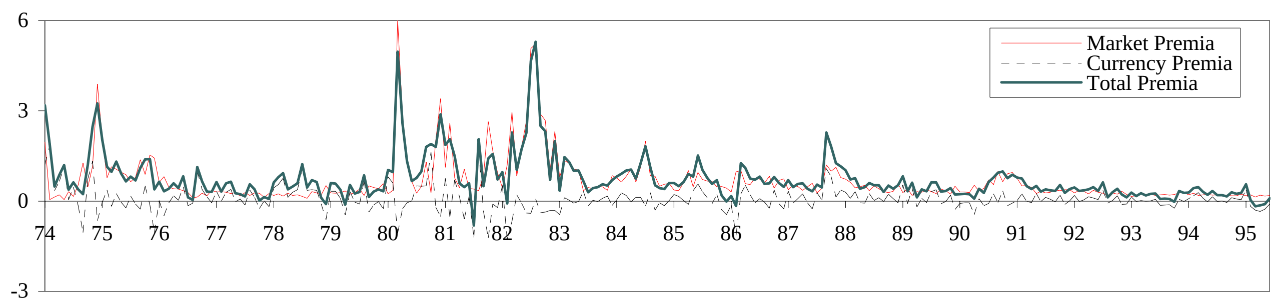


Figure 3f: Eurofranc Deposits

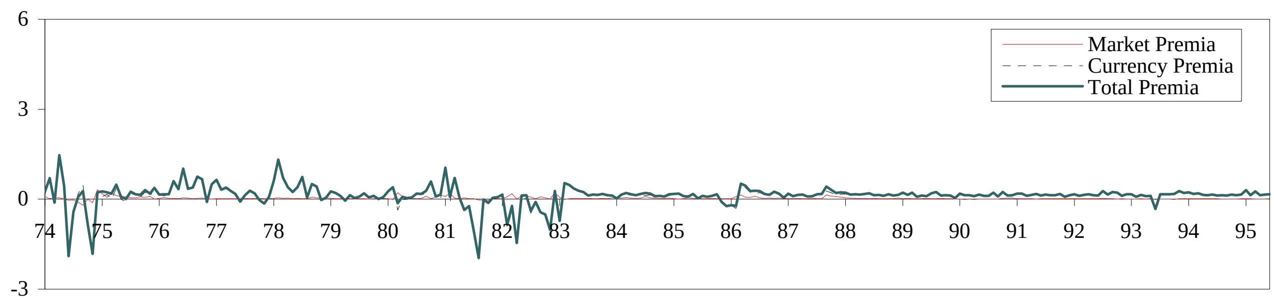


Figure 3g: Eurolira Deposits

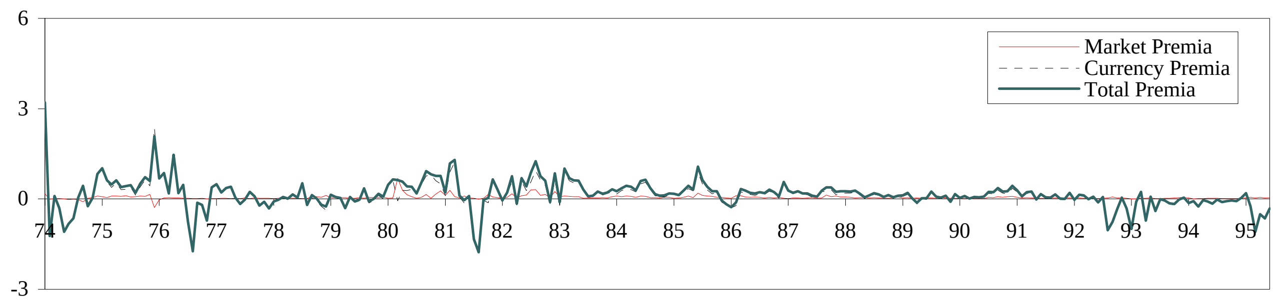


Figure 3h: Europound Deposits